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Comparing micro-evidence on rent sharing from two different econometric models[☆]

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ABSTRACT

The extent to which employers share rents with their employees is typically assessed by estimating the responsiveness of workers' wages on firms' ability to pay. This paper compares rent-sharing estimates using such a wage determination regression with estimates based on a productivity regression that relies on standard firm-level input and output data. Using a large matched firm-worker panel data sample for French manufacturing, we find that the respective industry distributions of the rent-sharing estimates are correlated and slightly overlap, but are significantly different on average. Precisely, if we only rely on the firm-level information, we obtain an average rent-sharing estimate of roughly 0.30 for the productivity regression and 0.17 for the wage determination regression. When we also take advantage of the worker-level information to control for unobserved worker ability in the model of wage determination, we find as expected a lower average value of 0.10.

1. Introduction

Contrary to the Walrasian labor market model, various non-competitive models predict a positive relationship between wages of comparable workers and the performance of their firms. Collective bargaining, optimal labor contract and search-theoretic models of the labor market share this theoretical conjecture, and consider different channels through which employer's ability to pay might affect wages.

We can view the wage determination equations specifying the expected positive wage-performance link as reduced-form models stemming from, or at least compatible with, such an underlying variety of theoretical structural models. Many empirical studies have estimated these reduced-form wage equations on firm data to test the rent-sharing hypothesis.[1] They have confirmed without exception that changes in firm performance feed through into changes in wages. In general, the estimated elasticities between wages and rents or profits per employee range between 0.05, even less, and 0.20, depending in particular on the quality of the instruments used to control for the endogeneity of profits. Following the seminal contribution of Abowd et al. (1999), more recent studies using matched employer-employee datasets, are able to include separately in the wage equations firm and worker effects that take into account the non-random sorting of high-ability (and thus high-wage) workers into high-profit firms. Compared to studies based on firm-level data only, these studies typically obtain, as expected, smaller estimates of wage-profit elasticities ranging from 0.01 to 0.10.[2]

Even more recently, a small set of productivity studies have extended the more standard productivity framework with imperfect competition in the product market to encompass two polar models of wage determi-

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[1] See in particular Barth et al. (2016) for the US; Abowd and Lemieux (1993) for Canada; Teal (1996) for Ghana; Van Reenen (1996) and Hildreth (1998) for the UK; Goos and Konings (2001) and Brock and Dobbelaere (2006) for Belgium; and Blanchflower et al. (1990), Nickell and Andrews (1983) and Hildreth and Oswald (1997) for a sample of European countries.

[2] See in particular Margolis and Salvanes (2001) for France and Norway; Kramarz (2003) and Fakhfakh and Fitzroy (2004) for France; Bronars and Famulari (2001) for the US; Arai (2003), Nekby (2003), Arai and Heyman (2009) and Carlsson et al. (2016) for Sweden; Bagger et al. (2014) for Denmark, Rycx and Tojerow (2004) and Du Caju et al. (2011) for Belgium; Guertzgen (2009) for Germany; Card et al. (2014) for Italy; and Cardoso and Portela (2009), Martins (2009) and Card et al. (2018) for Portugal.

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nation in imperfect labor markets.[3] These studies have also been able to provide estimates of the extent of rent sharing between firms and workers, and more specifically estimates of the corresponding wage-profit elasticities which are higher, in the [0.10-0.50] range.[4]

Our contribution to the empirical rent-sharing literature in this paper is to compare the rent-sharing estimates obtained in the case of French manufacturing for a large matched firm-worker panel data sample by relying on the wage determination and the productivity models. It is also to suggest potential explanations for the estimated discrepancies and to assess the advantages and shortcomings of both types of models.

The plan of the paper is as follows. Section 2 presents the two econometric models while Section 3 describes the data and explains the method of estimation. Section 4 compares and discusses estimates of the extent of rent sharing that we obtain from estimating the productivity and wage equations. Section 5 provides potential explanations of discrepancies between these estimates while Section 6 concludes.

2. Estimating rent sharing from two econometric models

We present in this Section the econometric reduced-form productivity and wage determination models as they have been usually specified in the literature and as we take them here to the data to better compare the estimates of extent of rent sharing they entail.

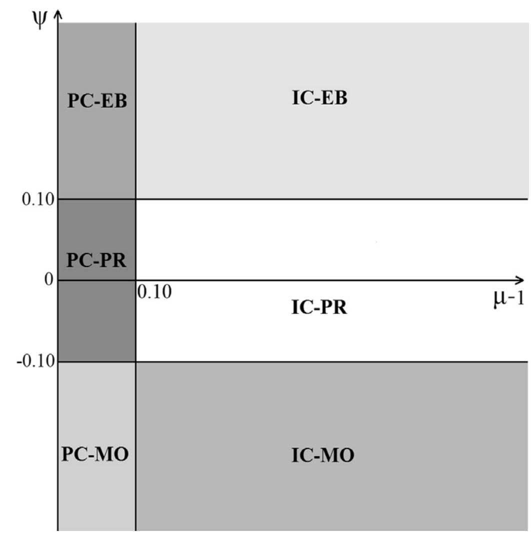
2.1. Reduced-form model of productivity

The specification of the reduced-form productivity equation we estimate is the following log-linear regression:

$$\begin{aligned} q_{it} = {} & \mu[s_{Nit}(n_{it} - k_{it}) + s_{Mit}(m_{it} - k_{it})] + \psi[s_{Nit}(k_{it} - n_{it})] \\ & + \lambda k_{it} + \omega_{it} + \alpha_i + \alpha_t + \epsilon_{it} \end{aligned} \quad (1)$$

where i is a firm subscript and t a year subscript. The variables q_{it}, n_{it}, m_{it} and k_{it} are respectively for each year the logarithms of output Q_{it}, labor N_{it}, material input M_{it} and capital K_{it}. s_{Nit} and s_{Mit} are for each year the shares of labor costs and material costs in total revenue. The parameters μ, ψ and λ are respectively the parameters of price-cost markup, joint product and labor market imperfections and elasticity of scale. ω_{it} is an index of "true" total factor productivity, or productivity for short, possibly observed by the firm at t when input choices are made, but unobserved to the econometrician. α_i is a firm-specific effect proxying for firm unobserved heterogeneity such as managerial ability differences, α_t is a year effect proxying for changes in firms' industrial environment, and ϵ_{it} is an idiosyncratic error term including non-predictable output shocks and potential measurement error in output and inputs.

As explained in Section 1 of the online supplementary material, we can distinguish six combinations or regimes of imperfect and "perfect or nearly perfect" competition in product and labor markets, which are based on the respective values of the price-cost mark-up and joint market imperfections parameters μ and ψ. We differentiate imperfect and nearly perfect product market settings on the basis of a price-cost markup μ higher than 1.1. Similarly, we separate the two settings of imperfect competition in the labor market, efficient bargaining and monopsony, from nearly perfect competition or right-to-manage bargaining in the labor market on the basis of a joint market imperfections parameter ψ respectively positive and higher than 0.1 or negative and smaller than -0.1. These threshold values, although conventional, are empirically reasonable. They also have the practical advantage of characterizing the different regimes as subsets of dimension 2 in the space of the two parameters μ and ψ (with $\mu \geq 1$), and they thus put the different regimes on a par when estimating their probability and testing that an industry or a selected group of firms belongs to a particular regime. The six regimes that we can thus consider are shown in Graph 1 in the two-dimensional space of the parameters μ and ψ .



Graph 1. Comparative analysis sample: IC-EB regime.
Notes: Product market settings: PC refers to perfect or "nearly perfect" competition and IC to imperfection competition, labor market settings: PR refers to perfect or "nearly perfect" competition or right-to-manage bargaining, EB to efficient bargaining and MO to monopsony. μ: price-cost markup, ψ: joint market imperfections parameter. PC-PR: $1 \leq \mu \leq 1.1$ and $-0.1 \leq \psi \leq 0.1$, PC-EB: $1 \leq \mu \leq 1.1$ and $\psi > 0.1$, PC-MO: $1 \leq \mu \leq 1.1$ and $\psi < -0.1$, IC-PR: $\mu > 1.1$ and $-0.1 \leq \psi \leq 0.1$, IC-EB: $\mu > 1.1$ and $\psi > 0.1$, IC-MO: $\mu > 1.1$ and $\psi < -0.1$.

More precisely, they are the following:

- $1 \leq \mu \leq 1.1$ and $-0.1 \leq \psi \leq 0.1$, or PC-PR, corresponding to perfect or "nearly perfect" competition in the product market, and perfect or "nearly perfect" competition or right-to-manage bargaining in the labor market.
- $1 \leq \mu \leq 1.1$ and $\psi > 0.1$, or PC-EB, corresponding to perfect or "nearly perfect" competition in the product market, and efficient bargaining in the labor market.
- $1 \leq \mu \leq 1.1$ and $\psi < -0.1$, or PC-MO, corresponding to perfect or "nearly perfect" competition in the product market, and monopsony in the labor market.
- $\mu > 1.1$ and $-0.1 \leq \psi \leq 0.1$, or IC-PR, corresponding to imperfect competition in the product market, and perfect or "nearly perfect" competition or right-to-manage bargaining in the labor market.
- $\mu > 1.1$ and $\psi > 0.1$, or IC-EB, corresponding to imperfect competition in the product market, and efficient bargaining in the labor market.
- $\mu > 1.1$ and $\psi < -0.1$, or IC-MO, corresponding to imperfect competition in the product market, and monopsony in the labor market.

Here, for the sake of comparison, we focus our analysis on the set of industries in which we expect that rent sharing is likely to prevail (IC-EB) on the basis of descriptive statistics as well as previous econometric studies where the estimates we found for the parameter of joint market

[3] This extension of the econometric productivity model to take into account imperfect labor markets has been developed in Dobbelaere and Mairesse (2013), after a first extension by Crépon et al. (2005), and following the revival of the empirical literature on productivity with imperfect product markets (Hall, 1988). Both extensions of econometric productivity analyses with imperfectly competitive product and labor markets find their historical roots in Marschak and Andrews (1944).

[4] Dobbelaere and Vancauteren (2014) use firm-level data for Belgium and the Netherlands, Dobbelaere et al. (2015) for France, Japan and the Netherlands, Dobbelaere et al. (2016) for Chile and France, Dobbelaere and Kiyota (2017) for Japan and Félix and Portugal (2017) for Portugal. Dobbelaere (2004), Abraham et al. (2009), Boulhol et al. (2011) and Ahsan and Mitra (2014) implement the extension of the econometric productivity model developed in Crépon et al. (2005).

imperfections ψ were positive. In this case, the parameter of main interest is the absolute extent of rent sharing sharing ϕ or equivalently the relative extent of rent sharing γ with $\gamma = \frac{\phi}{1-\phi}$ and $\phi = \frac{\gamma}{1+\gamma}$, which are estimated by:

$$\phi = \frac{\psi}{\mu + \psi} \quad \text{and} \quad \gamma = \frac{\psi}{\mu} \tag{2}$$

2.2. Reduced-form model of wage determination

The specification of the reduced-form wage determination equation is a log-log regression model that slightly differs depending on whether we estimate it on firm or matched firm-worker data. It can be written as Eq. (3) in the case of firm data:

$$w_{it} = \beta_0 + \beta_1 \overline{w}_{it} + \beta_2 (\pi_{it} - n_{it}) + \beta_3 (k_{it} - n_{it}) + \alpha_i + \alpha_t + \epsilon_{it} \tag{3}$$

and as Eq. (4) in the case of matched firm-worker data:

$$w_{j(i)t} = \beta_0 + \beta_1 \overline{w}_{it} + \beta_2 (\pi_{it} - n_{it}) + \beta_3 (k_{it} - n_{it}) + \alpha_{j(i)} + \alpha_i + \alpha_t + \epsilon_{it} \tag{4}$$

where i is a firm subscript and t a year subscript and $j(i)$ a subscript of worker j in firm i. The variables w_{it}, $w_{j(i)t}$, $\overline{w}_{it}$, π_{it}, k_{it} and n_{it} are respectively for each year the logarithms of the firm labor cost per worker or average wage W_{it}, the net earnings of worker j in firm i or the net wage $W_{j(i)t}$, the average workers' alternative wage or reservation wage $\overline{W}_{it}$, the firm profit or more generally economic rents Π_{it}, the firm capital K_{it}, and the firm number of employees N_{it} . The parameters β_1, β_2 and β_3 are the elasticities of wages with respect to the reservation wage, profit per employee and capital per employee, respectively. α_i is the firm effect, $\alpha_{j(i)}$ the worker-firm effect, α_t the year effect and ϵ_{it} an idiosyncratic error.

In the empirical literature, Eqs. (3) and (4) are commonly specified as a log-log regression in which case the relative extent of rent sharing γ is a varying parameter equal to the wage-profit elasticity β_2 multiplied by $Ratio_{it} = \frac{W_{it} \times N_{it}}{\Pi_{it}}$, the ratio of the firm wage bill to its profits.[5] For our purpose of comparison, we compare γ as estimated in the productivity regression (Eq. (1)) with its sample average values as estimated in the wage regressions (Eqs. (3) and (4)): $\gamma = \beta_2 \times \overline{Ratio}$ with $\overline{Ratio} = \text{mean}\left(\frac{W_{it} \times N_{it}}{\Pi_{it}}\right)$. The log specification of the wage regressions has econometric advantages, in particular by normalizing the wage distributions which are skewed to the right and display long right tails (Martins, 2007; Neal and Rosen, 2000). Adopting a log specification of both the productivity and wage equations is also appropriate in our case since it allows us to control for a potential source of discrepancy in the corresponding estimates of extent of rent sharing.

In practice, Eqs. (3) and (4) usually do not include the capital intensity variable and we also consider an estimation variant without it. We think it is preferable to take it here into account to control at least partly for differences in firms' labor skill composition and the possibility that rent sharing is relatively higher in capital-intensive firms. Skill-intensive firms will also be capital-intensive in industries where capital and skilled labor are complements (Bronars and Famulari, 2001; Duffy et al., 2004; Griliches, 1969). This implies that the wage-profit elasticity estimates will be upwardly biased for lack of suitable skill composition data but less so if we control for firm capital intensity. Note, in a related way, that such skill bias is likely to be largely controlled for when we take into account the non-random sorting of high-ability (hence high-wage) workers into high-productivity (hence high-profit) firms by relying on matched firm-worker data. We thus expect that the wage-profit elasticity will be less upwardly biased (hence smaller) when estimated in regression (4) than in regression (3).

[5] Note that if Eqs. (3) and (4) were specified as *linear* regressions, the parameter β_2 would simply be the parameter of relative extent of rent sharing, conditional on considering collective bargaining, among the various interpretative schemes, to be the main theoretical engine establishing a positive relationship between profitability and pay. This can be clearly seen from Eq. (20) in the online supplementary material. We elaborate on this in Section 5.

3. Data description and econometric identification

3.1. Comparative analysis sample and measurement of variables

We have constructed an unbalanced panel of French manufacturing firms over the period 1984-2001, based on confidential databases maintained by INSEE (the French "Institut National de la Statistique et des Etudes Economiques"): mainly firm accounting information from EAE ("Enquête Annuelle d'Entreprise"), supplemented by matched firm-worker data drawn from the DADS (the administrative database of "Déclarations Annuelles des Données Sociales"). We first trimmed the data to eliminate outliers and anomalies for our main variables: firm output and input growth rates, firm input shares in total revenue, firm average wages and profits, and worker net earnings. We then only kept firms with consecutive observations for at least four years and retained workers who remained in the same firms ("stayers"), worked twelve months per year and with consecutive observations for at least two years. We also retained the subset of 25 industries where we expect rent sharing to be predominant, chosen among the full set of 52 manufacturing industries defined on the basis of the 2- and 3-digit level of the French industrial classification ("Nomenclature économique de synthèse"). This subset amounts to 66% of the firms and 58% of employment in total manufacturing.

We thus end up with a matched firm-worker panel data sample, consisting at the firm level of 109,199 observations for 9,849 firms over the 18 years 1984–2001, with a median number of observations per firm of 11, and at the worker-firm level of 382,501 observations for 60,294 workers in the 9,849 firms, with a median number of 9 workers per firm.

The eleven variables involved in our regression analyses are defined and measured in the following way. Output (Q_{it}) is defined as current production deflated by the two-digit producer price index. Labor (N_{it}) refers to the average number of employees in each firm for each year. Material input (M_{it}) is defined as intermediate consumption deflated by the two-digit intermediate consumption price index. The capital stock (K_{it}) is measured by the gross book value of tangible fixed assets at the beginning of the year and adjusted for inflation. The shares of labor (s_{Nit}) and material input (s_{Mit}) are constructed by dividing respectively the firm total labor cost and intermediate consumption by the firm current production and by taking the average of these ratios over adjacent years. The firm average wage per worker (W_{it}) is computed as the wage bill divided by the average number of employees. The worker net wage ($W_{j(i)t}$) is the yearly net earnings of worker j in firm i and the number of workers ($N_{j(i)t}$) refers to the number of individual workers observed for firm i in the matched firm-worker sample. The firm profits (Π_{it}) is simply the widely used measure of gross operating profit computed as value added minus labor costs, smoothed over four or five years if possible from year $t-3$ or $t-4$ to current year t (taking advantage of the availability of three year pre-sample accounting firm observations when necessary). Such smoothing, often done in practice, is useful to control for the high volatility of profits. Finally, we rely on the matched firm-worker data to propose a measure of the average workers' alternative wage ($\overline{W}_{it}$) for the two wage regressions. In particular, we proxy the alternative wage by the 5th percentile of the workers' wage distribution but also consider an estimation variant in which it is measured by the 1st percentile.

Table 1 reports descriptive statistics for all our variables: mean, standard deviation, and first quartile, median and third quartile. The median number of employees is 49 and the mean number 123, while the median number of individual workers observed per firm is 9 and the mean number 21. The average yearly growth rate over the period 1984–2001 of firm output, number of employees, materials and capital are respectively 2.6%, 0.9%, 4.4% and 0.3%. The average shares of labor and materials in total revenue are of 33% and 49%. The median and mean are both of about 27,000 Euros for the workers' wage per year and respectively about 13,500 and 16,000 Euros for their net earnings, while

Table 1
Descriptive statistics: Comparative analysis sample, 1984-2001.

Variables	Mean	sd	Q_1	Q_2	Q_3	N
Real firm output growth rate (Δq_{it})	0.026	0.150	−0.058	0.022	0.108	96,508
Labor growth rate (Δn_{it})	0.009	0.126	−0.042	0.000	0.060	96,508
Materials growth rate (Δm_{it})	0.044	0.194	−0.061	0.039	0.145	96,508
Capital growth rate (Δk_{it})	0.003	0.154	−0.072	−0.018	0.067	96,508
Labor share in nominal output (s_{Nit})	0.328	0.136	0.231	0.314	0.407	109,199
Materials share in nominal output (s_{Mit})	0.494	0.150	0.401	0.502	0.599	109,199
Profits per employee $(\frac{\Pi}{N})_{it}$	19,392	24,491	6,790	12,678	23,309	109,199
Smoothed profits per employee $\widetilde{(\frac{\Pi}{N})}_{it}$	19,734	22,939	7,958	13,415	23,512	109,199
Firm average number of employees (N_{it})	123	255	32	49	116	109,199
Firm average wage per worker (W_{it})	27,381	7,612	21,944	26,667	31,907	109,199
Wage premium ($W_{it} - \overline{W}_{it}$)	8,103	6,283	3,662	7,023	11,372	109,199
Number of employees ($N_{j(i)t}$)	21	42	3	9	22	382,501
Average wage per worker ($W_{j(i)t}$)	15,919	8,882	10,807	13,690	18,046	382,501

Note: $\overline{W}_{it}$ is proxied by the 5th percentile value of the workers' wage distribution, $\widetilde{(\frac{\Pi}{N})}_{it}$ is defined as $\frac{1}{5}\sum_{k=t-4}^{t}(\frac{\Pi}{N})_{ik}$ if $(\frac{\Pi}{N})_{it-4}$ is not missing, otherwise equal to $\frac{1}{4}\sum_{k=t-3}^{t}(\frac{\Pi}{N})_{ik}$ (taking advantage of the availability of three year pre-sample accounting firm observations when necessary).

they amount to about 13,500 and 20,000 Euros per year for smoothed profits per employee.

3.2. Econometric identification and estimation

In the reduced-form productivity regression (Eq. (1)), we cannot assume that the input factor variables n_{it}, m_{it} and k_{it} are exogenous, even when we control for firm effects, and we cannot in general rely on ordinary least squares (OLS) estimation, even if we control for firm fixed effects by relying on the time dimension of the panel (i.e., the first-difference or within-firm dimension of the data). The crux of the identification problem inherent in estimating Eq. (1) is that a firm's choice of inputs (n_{it}, m_{it}, k_{it}) will likely depend on realized firm-specific productivity ω_{it}, which only the firm observes. Hence, we have to use an instrumental variable (IV) estimation method (as emphasized in the econometric production function literature since Marschak and Andrews, 1944; see also Griliches and Mairesse, 1998; Ackerberg et al., 2015). Similarly, we cannot assume that in the wage regressions (Eqs. (3) and (4)), the right-hand-side variables, in particular the profit-per-employee variable $(\pi_{it} - n_{it})$, are exogenous. Hence, they need to be instrumented. The endogeneity of profits is due to two sources of reverse causality. First, the wage-profit elasticity (the parameter β_2 in Eqs. (3) and (4)) might be underestimated due to the accounting relationship between wages and profits, implying that higher wages lead to lower profits. Second, theories of incentive pay and efficiency wages (Akerlof and Yellen, 1986; Shapiro and Stiglitz, 1984) predict that higher wages might lead to higher profits, which could generate an upward bias in wage-profit elasticities.

In order to get consistent estimates of the parameters in the productivity and wage equations, we apply the system generalized method of moments (SYS–GMM) estimation method, developed by Arellano and Bover (1995) and Blundell and Bond (1998), which is designed for panels with relatively small time and large cross-sectional dimensions, covariates that are not strictly exogenous, unobserved heterogeneity, heteroscedasticity and within-firm autocorrelation. This method extends the first-differenced GMM estimation method of Arellano and Bond (1991), by relying on a richer set of orthogonality conditions, which are obtained not only by using lagged variables in levels to instrument the equation written in first-differences, but also by using the lagged variables in first-differences to instrument the original equation in levels. Actually, to avoid instrument proliferation, we only exploit the orthogonality conditions entailing as instruments the 2- and 3-year lags of variables in levels and the 1-year lag of the first-differenced variables. We also use the two-step SYS–GMM estimator which is asymptotically more efficient than the one-step SYS–GMM estimator and robust to heteroscedasticity, and the finite-sample correction to the two-step covariance matrix developed by Windmeijer (2005).[6]

Data limitations precluded us from using exogenous firm demand shifters as a source of variation in input demands to obtain consistent estimates of the parameters in the reduced-form productivity equation (Eq. (1)). We follow a common instrumentation strategy in the literature, which is using lagged internal values. More explicitly, we use suitable past levels and differences of input factors as instruments for current inputs. This instrumentation strategy can be theoretically justified through adjustment costs generating dependence of current input levels on past realizations of productivity shocks (see Bond and Söderbom, 2005). Similarly, we lack exogenous firm demand shifters as a source of variation of profits that does not impact directly upon wages. Therefore, we follow common practice and use lagged values of firm profitability as instruments (see e.g. Blanchflower et al., 1996; Hildreth and Oswald, 1997), which in our case are suitable past levels and differences of the smoothed profits-per-employee variable.

We have chosen to restrict estimation of the wage-profit elasticity (β_2) in Eq. (4) to workers staying in the same firms over several years. More specifically, we control for both an unobserved firm fixed effect α_i and an unobserved worker within firm fixed effect $\alpha_{j(i)}$, or unobserved spell effect $\theta_s = \alpha_{j(i)} + \alpha_i$ for each unique worker-firm combination (Andrews et al., 2006). We do not separately control for unobserved firm fixed effects and unobserved worker fixed effects, and do not compute their empirical correlation, as is often done in matched employer-employee panel data analyses (see Abowd et al., 1999 and related references)[7]. This choice is grounded on a data reason and an econometric

[6] As the validity of GMM crucially hinges on the assumption that the instruments are exogenous, we use the Sargan and Hansen test statistics for the joint validity of the overidentifying restrictions. In addition to the Hansen test evaluating the entire set of overidentifying restrictions/instruments, we provide difference-in-Hansen statistics to test the validity of subsets of instruments. Details on testing for instrument exogeneity are provided in Section 3 in the online supplementary material.

[7] Following Abowd et al. (1999), the degree of assortative matching (or sorting) in the labor market is typically assessed by calculating the empirical corre-

one. The first is that separate identification of these types of unobserved fixed effects relies on workers who move between employers and that we can only trace the mobility of workers from one firm to another for a too small part of our sample.[8] The second is that the separate identification requires exogeneity of worker mobility, which is not likely to be the case or would need having information on the reasons for mobility and being able to implement some form of instrumentation (Gibbons and Katz, 1992; Goux and Maurin, 1999; Murphy and Topel, 1987).

4. Results of comparative analysis

We compare and discuss in this Section the industry-level estimates of relative and absolute extent of rent sharing $\widehat{\gamma}$ and $\widehat{\phi}$ that we obtain from the productivity and two wage regressions (Eqs. (1), (3) and (4), respectively), and that we label $\widehat{\gamma}_I^{prod}$, $\widehat{\gamma}_I^{wage,f}$ and $\widehat{\gamma}_I^{wage,w}$, and $\widehat{\phi}_I^{prod}$, $\widehat{\phi}_I^{wage,f}$ and $\widehat{\phi}_I^{wage,w}$, where subscript I varying from 1 to 25 stands for the different industries of our matched firm-worker panel. The standard errors of $\widehat{\gamma}_I^{prod}$ and $\widehat{\phi}_I^{prod}$ are computed using the Delta method (Wooldridge, 2002).[9] Details on estimates of output elasticities, wage-profit elasticities and other parameter estimates are relegated to Section 3 in the online supplementary material.

Table 2 reports these estimates ranked by increasing order of $\widehat{\gamma}_I^{prod}$ and $\widehat{\phi}_I^{prod}$. We see that while they vary between a minimum of about 0.10 and a maximum of about respectively 1.15 and 0.55, $\widehat{\gamma}_I^{wage,f}$ and $\widehat{\phi}_I^{wage,f}$ vary between negative or zero values in three industries (−0.10, −0.05 and −0.00) and a maximum of about respectively 0.85 and 0.45, and $\widehat{\gamma}_I^{wage,w}$ and $\widehat{\phi}_I^{wage,w}$ vary between negative and a value of less than 0.05 in ten industries and a maximum of about respectively 0.85 and 0.45. A visual representation of the sampling distribution of the rent-sharing estimates is given in Graph 2. These box plots summarize well the average overall picture by abstracting from “outlier” industry estimates. They show that the three sets of estimates differ clearly in average, although they tend to overlap slightly. In the case of the productivity regression, we find median industry estimates of relative and absolute extent of rent sharing of respectively 0.41 and 0.29 to be compared to 0.19 and 0.16 in the case of the firm-level wage regression and 0.09 and 0.08 in the case of the worker-firm wage regression. The evidence is roughly that of a difference of 0.1 between the estimates from the two wage equations, and of 0.2 or 0.3 between them and the ones from the productivity equation.

Beyond their overall differences, Tables 3 and 4 provide a comprehensive view of the similarity of the sampling distributions of the three sets of industry estimates of rent sharing. Table 3 shows that the correlations between $\widehat{\gamma}_I^{wage,f}$ and $\widehat{\gamma}_I^{wage,w}$ and between $\widehat{\phi}_I^{wage,f}$ and $\widehat{\phi}_I^{wage,w}$ are high and statistically significant at the 10% level of confidence. These correlations are about 0.35 for the usual Spearman's rank correlation coefficients and respectively 0.28 and 0.41 for the robust “bi-weight mid-correlation” or Wilcox (2012) coefficients. They are also sizeable for the correlations between $\widehat{\gamma}_I^{prod}$ and $\widehat{\gamma}_I^{wage,f}$ and between $\widehat{\phi}_I^{prod}$ and $\widehat{\phi}_I^{wage,f}$. They are about 0.25 for both the Spearman and Wilcox coefficients and statistically significant at the 5% or 10% level of confidence for the latter. However, the corresponding correlations between $\widehat{\gamma}_I^{prod}$ and $\widehat{\gamma}_I^{wage,w}$ and between $\widehat{\phi}_I^{prod}$ and $\widehat{\phi}_I^{wage,w}$ are small, even negative of about −0.05, for the Spearman coefficients, and they are positive and sizeable of about 0.30 and 0.20, if not statistically significant, for the Wilcox coefficients. This is a mixed picture, but not a bad one if we take into consideration that these correlations are computed for distributions of 25 estimates only, and they concern a subset of very diverse and heterogeneous industries.

Table 4 gives the mean, first quartile Q_1, median Q_2 and third quartile Q_3 of the three sets of industry estimates of rent sharing. It shows that the differences between them in the mid-range of their distribution, from Q_1 to Q_3, are roughly the same as already mentioned for the median: of 0.1 between the estimates from the two wage equations, and of 0.2 or 0.3 between them and the ones for the productivity equation. The first quartile value of $\widehat{\gamma}_I^{wage,w}$ is the only noteworthy exception to such nearly constant shift. By itself, it suggests that a common omitted variable misspecification, namely workers' skills, could be a potential explanation to the extent that it would affect the three sets of rent-sharing estimates differentially. This is what we try to substantiate among other a priori sources of discrepancies in the next Section.

5. Potential sources of discrepancies between rent-sharing estimates

We can a priori distinguish three large categories of reasons or sources of discrepancies that we find between the distributions of the industry rent-sharing parameters as estimated on the basis of the productivity and wage determination regressions (Eqs. (1), (3) and (4)).

A first category is economic specification errors, which involve omitted relevant variables as well as poor measurement of included available variables. An important case, for the wage regressions, particularly for regression (3), is the omission of a variable or group of variables of workers' skills because of the lack of suitable skill composition data at the firm level. We expect that rent sharing is higher in skill-intensive firms, and hence that wage-profit elasticities will be upwardly biased in the absence of skill variables in the two wage regressions. Actually, the omission of skill variables is the most likely source of the smaller estimates of rent sharing found with regression (4) than with regression (3). As already explained, the specification of these two regressions is basically the same, their main difference being that they are estimated at the worker-firm level and firm-level of our matched firm-worker panel data sample. At the worker-firm level, we can expect that the worker-firm effect is positively correlated with the worker's skills. We know in fact that the assortative matching of firms and workers is non-random, and that high-skilled (and thus high-wage) workers tend to be selected into high-productive (and thus high-profit) firms (Abowd et al., 1999; Card et al., 2018).

As already mentioned, the introduction of the capital-per-employee variable in the two wage regressions is largely to proxy for the omission of skill variables. As capital-intensive firms will also be skill-intensive in industries where capital and skilled labor are complements, we expect that wage-profit elasticity estimates will be less upwardly biased for lack of skill variables when the capital-per-employee variable is included in the wage regressions. This is indeed confirmed if we estimate them without this variable. We find that the mean (or median) of the industry-level absolute extent of rent-sharing estimates $\widehat{\phi}_I^{wage,f}$ and $\widehat{\phi}_I^{wage,w}$ increase respectively from 0.17 to 0.25 (0.16 to 0.23) and from 0.10 to 0.14 (0.08 to 0.15), that is, in average by about respectively 50% and 40% (by about respectively 45% and 90%). In terms of comparison with the productivity regression industry estimates, when we do not include the capital-per-employee variable in the two wage regressions, the differences between the mean (or median) industry-level absolute extent of rent-sharing estimates $\widehat{\phi}_I^{prod}$ and respectively $\widehat{\phi}_I^{wage,f}$ and $\widehat{\phi}_I^{wage,w}$ decrease in average by about respectively 65% and 20% (by about respectively 55% and 35%).

lation between worker and firm fixed effects (see e.g. Goux and Maurin (1999), Abowd et al. (2004) and Abowd et al. (2009) for France and the US; Gruetter and Lalive (2009) for Austria, Andrews et al. (2008) for Germany; and Sørensen and Vejlin (2013) for Denmark). Woodcock (2015) provides evidence on assortative matching for the US by directly controlling for an interaction effect between the worker and the firm.

[8] Andrews et al. (2008, 2012) show that the downward bias in the estimated correlation between worker and firm fixed effects is larger when there are fewer movers in the data (labelled “limited mobility bias”) using German data.

[9] Dropping subscripts, $(\sigma_{\widehat{\gamma}})^2 = \left(\frac{s_M}{s_N+s_M-1}\right)^2 \frac{(\widehat{\varepsilon}_M^Q)^2\left(\sigma_{\widehat{\varepsilon}_N^Q}\right)^2 - 2\widehat{\varepsilon}_N^Q\widehat{\varepsilon}_M^Q\left(\sigma_{\widehat{\varepsilon}_N^Q,\widehat{\varepsilon}_M^Q}\right) + (\widehat{\varepsilon}_N^Q)^2\left(\sigma_{\widehat{\varepsilon}_M^Q}\right)^2}{(\widehat{\varepsilon}_M^Q)^4}$ and $\left(\sigma_{\widehat{\phi}}\right)^2 = \frac{(\sigma_{\widehat{\gamma}})^2}{(1+\widehat{\gamma})^4}$ where $\widehat{\varepsilon}_N^Q$ and $\widehat{\varepsilon}_M^Q$ are the estimated output elasticities with respect to labor and material input, respectively.

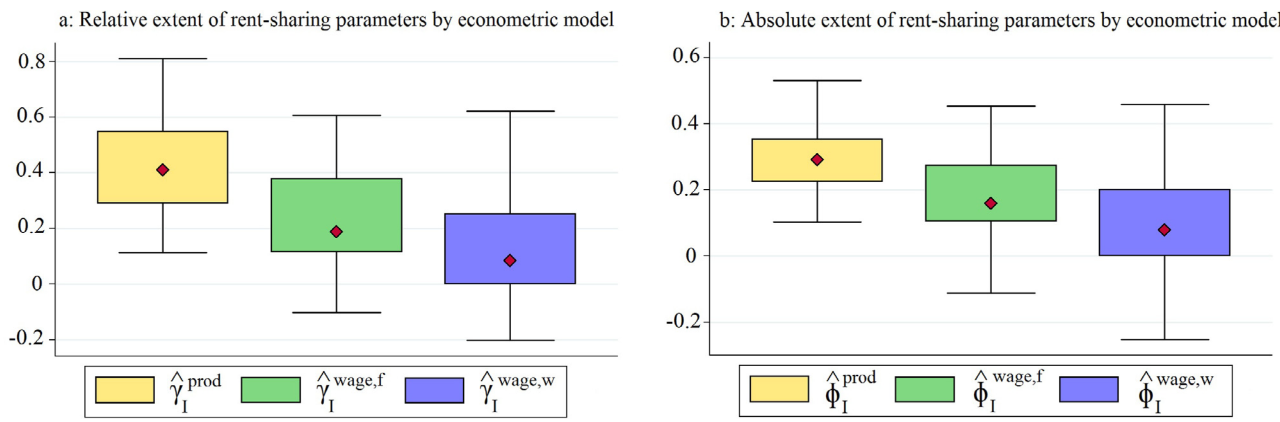
Table 2
Reduced-form models of productivity and wage determination: Industry-specific relative ($\hat{\gamma}_I^{prod}, \hat{\gamma}_I^{wage,f}, \hat{\gamma}_I^{wage,w}$) and absolute ($\hat{\phi}_I^{prod}, \hat{\phi}_I^{wage,f}, \hat{\phi}_I^{wage,w}$) extent of rent-sharing parameters.

		Reduced-form model of productivity		Reduced-form model of wage determination			
				Dep. var.: Firm wage w_{it}		Dep. var.: Worker wage $w_{j(i)t}$	
Ind. I	Name	$\hat{\gamma}_I^{prod}$	$\hat{\phi}_I^{prod}$	$\hat{\gamma}_I^{wage,f}$	$\hat{\phi}_I^{wage,f}$	$\hat{\gamma}_I^{wage,w}$	$\hat{\phi}_I^{wage,w}$
17	Articles of paper and paperboard	0.115 (0.196)	0.103 (0.158)	−0.037 (0.056)	−0.039 (0.060)	0.058 (0.050)	0.054 (0.045)
19	Plastic products	0.141 (0.138)	0.124 (0.106)	−0.101 (0.135)	−0.112 (0.167)	0.199 (0.089)	0.166 (0.062)
8	Metal products for construction	0.210 (0.372)	0.173 (0.254)	0.517 (0.608)	0.341 (0.264)	0.849 (0.440)	0.459 (0.129)
13	Earthenware products and construction material	0.231 (0.129)	0.188 (0.085)	0.187 (0.088)	0.158 (0.062)	0.124 (0.062)	0.111 (0.049)
4	Publishing, (re)printing	0.255 (0.157)	0.203 (0.099)	0.210 (0.203)	0.174 (0.139)	−0.170 (0.113)	−0.204 (0.164)
12	Mining of metal ores, other mining n.e.c.	0.286 (0.115)	0.223 (0.069)	0.106 (0.142)	0.096 (0.116)	0.086 (0.072)	0.079 (0.061)
21	Production of non-ferrous metals	0.291 (0.279)	0.225 (0.168)	0.041 (0.093)	0.039 (0.086)	0.007 (0.058)	0.007 (0.057)
16	Pulp, paper and paperboard	0.316 (0.354)	0.240 (0.204)	0.137 (0.134)	0.120 (0.104)	0.252 (0.094)	0.202 (0.060)
11	Medical and surgical equipment and orthopaedic appliances	0.317 (0.578)	0.241 (0.333)	0.387 (0.472)	0.279 (0.245)	0.333 (0.267)	0.250 (0.150)
20	Basic iron and steel	0.319 (0.220)	0.242 (0.127)	0.004 (0.225)	0.004 (0.224)	−0.053 (0.052)	−0.056 (0.058)
1	Other food products	0.369 (0.150)	0.269 (0.080)	0.224 (0.151)	0.183 (0.101)	0.089 (0.085)	0.082 (0.072)
3	Leather goods and footwear	0.373 (0.278)	0.272 (0.147)	0.378 (0.122)	0.274 (0.064)	0.142 (0.097)	0.124 (0.075)
18	Rubber products	0.411 (0.341)	0.291 (0.171)	0.279 (0.118)	0.218 (0.072)	0.001 (0.045)	0.001 (0.045)
6	Shipbuilding, construction of railway rolling stock, bicycles, motorcycles, transport equipment n.e.c.	0.421 (0.519)	0.296 (0.257)	0.607 (0.160)	0.378 (0.062)	0.297 (0.149)	0.229 (0.089)
7	Aircraft and spacecraft	0.469 (0.577)	0.319 (0.267)	0.116 (0.504)	0.104 (0.405)	−0.202 (0.105)	−0.253 (0.165)
5	Furniture	0.478 (0.231)	0.323 (0.106)	0.189 (0.176)	0.159 (0.125)	0.249 (0.135)	0.199 (0.087)
25	Electronics	0.479 (0.300)	0.324 (0.137)	0.425 (0.348)	0.298 (0.171)	0.710 (0.191)	0.415 (0.065)
22	Ironware	0.482 (0.220)	0.325 (0.100)	0.171 (0.122)	0.146 (0.089)	0.013 (0.085)	0.013 (0.083)
10	Other special purpose machinery	0.550 (0.405)	0.355 (0.169)	0.200 (0.326)	0.167 (0.226)	−0.006 (0.172)	−0.006 (0.174)
9	Ferruginous and steam boilers	0.599 (0.292)	0.374 (0.114)	0.462 (0.547)	0.316 (0.256)	−0.060 (0.196)	−0.064 (0.222)
15	Knitted and crocheted fabrics	0.657 (0.321)	0.397 (0.117)	0.284 (0.287)	0.221 (0.174)	−0.054 (0.149)	−0.057 (0.167)
24	Metal fabrication	0.685 (0.140)	0.406 (0.049)	0.124 (0.076)	0.110 (0.060)	0.034 (0.073)	0.033 (0.068)
14	Spinning and weaving	0.809 (0.256)	0.447 (0.078)	0.124 (0.158)	0.110 (0.125)	0.290 (0.113)	0.225 (0.068)
23	Industrial service to metal products	0.810 (0.147)	0.447 (0.045)	−0.005 (0.176)	−0.005 (0.178)	0.085 (0.136)	0.078 (0.116)
2	Clothing and skin goods	1.130 (0.223)	0.531 (0.049)	0.831 (0.215)	0.454 (0.064)	0.622 (0.190)	0.383 (0.072)

Notes: The standard errors in parentheses measure the dispersion of the rent-sharing parameters at the level of firms making up the industry. The standard errors of the average rent-sharing estimates are obtained by dividing the reported standard errors by the square root of the number of industry observations. $\hat{\phi}_I^{prod} = \frac{\hat{\gamma}_I^{prod}}{1+\hat{\gamma}_I^{prod}}$, $\hat{\gamma}_I^{wage,f} = \left(\hat{\beta}_2\right)_I^{wage,f} \times \left(\overline{Ratio}\right)_I$ with $\hat{\beta}_2$ the estimated wage-profit elasticity and $\overline{Ratio}$=mean$\left(\frac{W_{it} \times N_{it}}{\Pi_{it}}\right)$, $\hat{\phi}_I^{wage,f} = \frac{\hat{\gamma}_I^{wage,f}}{1+\hat{\gamma}_I^{wage,f}}$. Similar formulas apply if the dependent variable is the worker wage $w_{j(i)t}$. Industries are ranked according to $\hat{\gamma}_I^{prod}$.

A second category of potential sources of discrepancies between our reduced-form regression estimates of extent of rent sharing, which is closely related to the first category, concerns estimation methods, in particular the instrumentation strategy. Estimation of the productivity and the wage determination equations is based on exploiting different moment conditions for identification. Since data limitations precluded us from relying on external a priori exogenous instruments, we followed, as we explained, the common method of using past differences and levels of the regression variables themselves to construct such moment conditions. Since these variables are rather similar for the productivity and wage regressions, and hence the moment conditions based on their lagged differenced and level values are close, and since we have been attentive to check for their validity, it seems actually unlikely that estimation methods could be a significant source of discrepancies between our rent-sharing estimates.

A third category relates to different underlying theoretical structural models, which are themselves specifically or loosely related to various interpretative schemes. As stressed from the outset, both the econometric productivity and wage determination regressions are reduced-form models. Given that our main interest is in the assessment and comparison of firm-worker rent-sharing parameters, the reduced-form regressions provide satisfactory results if the first and second categories of discrepancies remain unimportant, that is, if they are appropriately specified and estimated. At the same time, we cannot and do not analyze the theoretical structural models which are possibly underlying these reduced-form models and are at least compatible with them. A consequence is that we are not able to decide between different interpretative schemes of our results, nor are we in a position to pin down the economic mechanisms and channels of rent sharing between firms and their workers.



Graph 2. Relative and absolute extent of rent sharing parameters by econometric model.
Notes: The box plots provide a summary of the sampling distributions of industry estimates of rent sharing. The upper and lower limits of the boxes represent the first and third quartiles of extent of rent-sharing parameter estimates while the median is represented by the diamond. Subscript "prod" denotes rent-sharing estimates obtained from the reduced-form model of productivity. Subscript "wage,f" denotes rent-sharing estimates obtained from the reduced-form model of wage determination using the firm average wage as the dependent variable. Subscript "wage,w" denotes rent-sharing estimates obtained from the reduced-form model of wage determination using the worker's wage as the dependent variable.

Table 3
Correlation of industry-specific relative and absolute extent of rent-sharing parameters across the reduced-form productivity model and the reduced-form model of wage determination.

Relative extent of rent sharing	$\hat{\gamma}_I^{prod}$	$\hat{\gamma}_I^{wage,f}$	$\hat{\gamma}_I^{wage,w}$
Model of productivity: $\hat{\gamma}_I^{prod}$	1.000 [1.000]		
Model of wage determination, dep. var. = w_{it} : $\hat{\gamma}_I^{wage,f}$	0.258 [0.253*]	1.000 [1.000]	
Model of wage determination, dep. var. = $w_{j(i)t}$: $\hat{\gamma}_I^{wage,w}$	−0.074 [0.302]	0.350*[0.278*]	1.000 [1.000]
Absolute extent of rent sharing	$\hat{\phi}_I^{prod}$	$\hat{\phi}_I^{wage,f}$	$\hat{\phi}_I^{wage,w}$
Model of productivity: $\hat{\phi}_I^{prod}$	1.000 [1.000]		
Model of wage determination, dep. var. = w_{it} : $\hat{\phi}_I^{wage,f}$	0.258 [0.266**]	1.000 [1.000]	
Model of wage determination, dep. var. = $w_{j(i)t}$: $\hat{\phi}_I^{wage,w}$	−0.074 [0.198]	0.350*[0.410*]	1.000 [1.000]

Notes: $\hat{\gamma}_I^{wage,f} = \left(\hat{\beta}_2\right)_I^{wage,f} \times \left(\overline{Ratio}\right)_I$, with $\hat{\beta}_2$ the estimated wage-profit elasticity and $\overline{Ratio}$ = mean$\left(\frac{W_{it} \times N_{it}}{\Pi_{it}}\right)$, $\hat{\phi}_I^{wage,f} = \frac{\hat{\gamma}_I^{wage,f}}{1+\hat{\gamma}_I^{wage,f}}$. Similar formulas apply if the dependent variable is the worker wage $w_{j(i)t}$. $\hat{\phi}_I^{prod} = \frac{\hat{\gamma}_I^{prod}}{1+\hat{\gamma}_I^{prod}}$. Spearman's rank correlation is reported. Wilcox' robust correlation is reported in square brackets. **significant at 5%, *significant at 10%.

Table 4
Comparison of the distribution of relative and absolute extent of rent-sharing parameters across the reduced-form productivity model and the reduced-form model of wage determination.

Reduced-form econometric model		Mean	Q_1	Q_2	Q_3
	Relative extent of rent sharing				
Model of productivity	$\hat{\gamma}_I^{prod}$	0.448	0.291	0.411	0.550
Model of wage determination, dep. var. = w_{it}	$\hat{\gamma}_I^{wage,f}$	0.234	0.116	0.189	0.378
Model of wage determination, dep. var. = $w_{j(i)t}$	$\hat{\gamma}_I^{wage,w}$	0.156	0.001	0.086	0.252
	Absolute extent of rent sharing				
Model of productivity	$\hat{\phi}_I^{prod}$	0.293	0.225	0.291	0.355
Model of wage determination, dep. var. = w_{it}	$\hat{\phi}_I^{wage,f}$	0.168	0.104	0.159	0.274
Model of wage determination, dep. var. = $w_{j(i)t}$	$\hat{\phi}_I^{wage,w}$	0.099	0.001	0.079	0.202

Notes: $\hat{\gamma}_I^{wage,f} = \left(\hat{\beta}_2\right)_I^{wage,f} \times \left(\overline{Ratio}\right)_I$, with $\hat{\beta}_2$ the estimated wage-profit elasticity and $\overline{Ratio}$ = mean$\left(\frac{W_{it} \times N_{it}}{\Pi_{it}}\right)$, $\hat{\phi}_I^{wage,f} = \frac{\hat{\gamma}_I^{wage,f}}{1+\hat{\gamma}_I^{wage,f}}$. Similar formulas apply if the dependent variable is the worker wage $w_{j(i)t}$. $\hat{\phi}_I^{prod} = \frac{\hat{\gamma}_I^{prod}}{1+\hat{\gamma}_I^{prod}}$.

In the online supplementary material, we present one theoretical structural model behind the reduced-form productivity regression, in which case the interpretative scheme is collective bargaining. We also present three potential theoretical structural models which can substantiate the expected positive pay-performance link of the wage determination regressions: collective bargaining models, an optimal labor contract model and a search-theoretic model of the labor market.[10] In collective bargaining models, the existence and strength of workers' bargaining power, which can correspond to different practices, institutionalized or not, is central.[11] In optimal contract models in which both workers and firms are risk-averse, the pay-performance link depends on the ratio of both parties' relative risk aversion parameters.[12] In two-sided search models with wage posting, the main source of rent sharing is competition between firms to attract workers. Firms have an incentive to hire more workers, thereby reducing search costs.[13] This incentive is particularly pronounced for higher-productivity firms because they face larger opportunity costs of search.

Although the three structural models can be developed analytically, their econometric analyses, in particular so that they could be compared together as well as with a structural productivity model with imperfections in product and labor markets, would be a formidable endeavor, if only because of multiple data constraints. To give two examples, one could estimate the coefficient of relative risk aversion from data on labor supply, but this would entail estimating wage and income elasticities, which could be done on the condition of having data on exogenous variation in unearned income and wages due to tax changes or lottery winnings. Similarly, one could attempt to quantify the extent to which search costs may drive a positive pay-performance link on the condition of having data on differential hiring activities across firms.

6. Conclusion

The basic objective of this paper is to compare as closely as possible rent-sharing estimates based on a reduced-form wage determination model adopted in a large empirical literature on firm-worker rent-sharing with rent-sharing estimates based on a reduced-form productivity model developed more recently, which we consider as largely complementary and which we think should provide reconcilable estimates. Grounded on a large matched firm-worker panel data sample, our main finding is that industry distributions of rent-sharing estimates are well correlated and overlap, but are nonetheless significantly different on average. Precisely, looking at the average industry estimates of the extent of rent-sharing parameter, we obtain an estimate of roughly 0.30 for the productivity regression and 0.17 for the wage determination regression if we rely only on firm-level information. If we also take advantage of the worker-level information to control for unobserved worker ability in the model of wage determination, thereby accounting for non-random sorting of high-ability (and thus high-wage) workers into high-profit firms, we find as expected a lower average value of 0.10.

There are a priori three large categories of reasons or potential sources of discrepancies that we find between the three types of estimates: economic specification errors, which involve omitted relevant or poorly measured variables; estimation methods, particularly instrumentation strategy; and different underlying theoretical structural models related to a variety of interpretative schemes. The idea of addressing all these potential sources of discrepancies in an encompassing model would be a formidable challenge, if only because of specific data requirements.

Renouncing to consider it, we can think of two interesting routes for future research. The first is to analyze and test separately the potential sources of discrepancies between the econometric reduced-form productivity and wage determination models, for example, by trying to take explicitly into account different workers' skills and by considering, more generally, that heterogeneity of firms and workers, of markets and industries is likely to be a dominant driving source of the discrepancies in our present estimates. The latter could be investigated on various sets of specific detailed datasets, corresponding to different periods, countries, industries, labor and product markets.

A complementary route of research is to empirically and specifically relate the reduced-form productivity and wage regressions to their underlying structural models by econometrically specifying and testing them in a multi-equation framework. There already exist many attempts in such direction (e.g. Bughin, 1993; Bughin, 1996; Forlani et al., 2016; Jaumandreu and Mairesse, 2010; Peters et al., 2017), but they still raise numerous difficulties. Actually, in the case of the reduced-form productivity regression, this endeavor brings us back to the paradigm of Marschak and Andrews (1944). As such, it involves framing a structural model composed of a production function, a demand function, a pricing rule, cost share equations for variable input factors, potentially taking into account some type of worker heterogeneity and separate wage equations for different types of workers.

Supplementary material

Supplementary material associated with this article can be found, in the online version, at doi:10.1016/j.labeco.2018.02.009.

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[10] Note that an expected positive pay-performance link can be derived from at least two other models. One is a modified version of the competitive labor market model with temporary frictions and a positively-sloped labor supply schedule (see Blanchflower et al., 1996). The other is the efficiency wage model in which increased productivity arises from reduced shirking and thus generates a positive wage-profitability correlation (Shapiro and Stiglitz, 1984).

[11] In these models, the pay-performance relationship depends on the relative strengths of the bargaining parties (see Eq. (20) in the online supplementary material).

[12] See Eq. (27) in the online supplementary material.

[13] See the combination of Eqs. (33) and (34) in the online supplementary material.

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